Downward Nominal Wage Rigidity: A Critical Assessment and some new evidence for Canada.[1]

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Abstract

In this paper, we review the existing empirical literature on the extent and consequences of downward nominal wage rigidities (DNWR). We conclude that recent studies, based on longitudinal micro data, provide compelling evidence that DNWR are an important labour market phenomena. However, it is less clear that DNWR have significant consequences for aggregate wage and employment (or unemployment) determination. One key empirical implication of DNWR is that, in response to a given negative shock, the real wage should decline less in periods of lower than higher inflation. Using a new Canadian wage series (based on individual data files from the Survey of Consumer Finances) for the period 1981 to 1997, we test this implication by estimating "real wage Phillips curves" that link the unemployment rate to the change in real wages. While the aggregate results are indecisive, the results based on cross-provincial variation indicate that the slope of this real wage Phillips curve has remained constant over time. These findings suggest that DNWR did not have a significant impact on aggregate wage and employment determination in Canada during the post-1991 period of low inflation.

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1 Introduction

Labour market observers have long suspected that, for a variety of reasons, employers are unwilling to reduce the nominal wages paid to their workers even when they experience severe financial difficulties.[1] Starting with Keynes' General Theory, these presumed downward nominal wage rigidities (hereafter, DNWR) have played a prominent role in many models of the labour market and the macroeconomy. One of Keynes' conjecture was that in a period of deflation like the Great Depression of the 1930s, DNWR resulted in higher real wages which made the Depression longer and deeper.

There has been renewed interest in DNWR over the last decade for several reasons. From a research perspective, the availability of rich longitudinal micro data sets has enabled researchers to formally test for the existence of DNWR. From an economic policy perspective, DNWR have become potentially more relevant to the conduct of economic policy as several countries have experienced very low rates of inflation in the 1990s. One argument which is closely related to Keynes' conjecture is that when inflation is very low, DNWR may prevent real wages to fall by as much as they should when the economy experiences negative shocks. For instance, Fortin (1996) uses this argument to explain why the recession of the 1990s was much longer and deeper in Canada, where CPI inflation averaged 1.4 percent from 1992 to 1997, than in the United States where CPI inflation averaged 2.9 percent during the same period.

The objective of this paper is two-fold. Our first goal is to critically review the existing literature on the extent and consequences of DNWR. From this review we conclude that recent studies, mostly based on U.S. longitudinal micro data, provide compelling evidence that DNWR are an important labour market phenomena. The main finding from this literature is that there is a sharp concentration of nominal wage changes at zero. The answer to the question of whether DNWR do in fact exist is a clear yes.

It is much less clear from the literature, however, that DNWR have significant consequences for aggregate wage and employment (or unemployment) determination. The second goal of the paper is, therefore, to take a new look at the effect of DNWR on wage and employment determination in Canada in periods of low inflation.

One reason why little research has been conducted on this topic in Canada is that limited wage data are available here relative to the United States. This explains why researchers such as Fortin (1996) and Craw-

[1] See Bewley (2000) for recent evidence.

ford and Harrison (1997) have used wage settlement data from collective agreement to look at the extent and consequences of DNWR in Canada. Unfortunately, these data from large firms in the unionized sector may not be very representative of the whole Canadian labour market.

To overcome these data shortcomings, we first develop a new wage series based on individual data files from Statistics Canada's Survey of Consumer Finance (SCF) for the period 1981 to 1997. This new wage series has several important advantages over what was previously available. First, it is based on a representative survey that can also be used to compute separate wage series by province, industry, etc. Second, it is possible to adjust wages for secular or business cycle changes in the composition of the workforce since detailed information is available on human capital (age, education, etc.) and job characteristics (industry, occupation, seniority, etc.) in this survey. This is an important issue since existing studies such as Solon, Barsky and Parker (1994) suggest that changes in the composition of the workforce tend to understate the cyclicality of wages over the business cycle.

Controlling for changes in the composition of the workforce is particularly important in the context of the impact of DNWR which are believed to apply only to workers who remain with the same employer. In a recession, aggregate wages may incorrectly look downward rigid if workers who lose their jobs earn systematically less than workers who keep them. This composition effect biases up aggregate wage changes which could mask real wage declines among workers who remain employed.

We then use this new wage series to analyze the relationship between real wage changes and economic conditions. One key empirical implication of DNWR is that, in response to a given negative shock, the real wage should decline less in periods of lower than higher inflation, because DNWR are not likely to bind in the former case than in the latter. We test this implication by estimating "real wage Phillips curves" that link the unemployment rate to the change in real wages. If DNWR prevents real wages from adjusting (downward) in periods of low inflation, the Phillips curve should be *flatter* in periods of lower inflation than in periods of higher inflation.

We use several empirical strategies to test this hypothesis that the Phillips curve became flatter when inflation dropped below 2 percent in the 1990s. First, we analyze the aggregate time series behavior of real wages and find that it is partly consistent with this hypothesis. Until 1992 (when the rate of inflation dropped "permanently" below two percent) there was a negative and statistically significant relationship between the unemployment rate and real wage changes. This relationship no longer holds since 1992, suggesting that real wages did not fall as much as they should have in the depth of the

1990s recession. One concern with these time series results, however, is that other unmodelled factors such as supply shocks or changes in the formation of expectations may also have changed during this period. Furthermore, the relationship between real wage changes and the unemployment rate is estimated imprecisely in the 1990s, because of small sample sizes.

Our second empirical strategy relies on variation both across time and provinces in economic conditions to identify potential changes in the relationship between unemployment rates and changes in real wages. Since different provinces are subject to different shocks at different points in time, it is possible, in principle, to identify the connection between (provincial) wage changes and (provincial) unemployment rates, while controlling for national wide factors using unrestricted year effects. Consistent with our expectations, we find that provinces that experience an increase in relative unemployment rates tend to experience a decline in relative wage growth. However, we do not find that this relationship has changed over time. In other words, these "provincial Phillips curves" did not become flatter in the years of very low inflation than they were earlier.

Finally, we exploit the richness of the SCF data to understand better the cyclical behavior of real wages in Canada from 1981 to 1997. We find that during both of the recession of 1981 83 and 1990 92, the real wages of older and more senior workers remained relatively constant. Most of the decline in real wages was concentrated among young workers and other workers having just started a new job. Irrespective of the inflation rate, new entrants seem to bear a disproportional share of the real wage adjustments over the business cycle. This may explain why DNWR, which most likely bind for older and more senior workers, seem to only have a modest impact on aggregate wages and employment.

The plan of the paper is as follows. In Section II, we present a critical assessment of the existing literature and highlight the major knowledge gaps on the effect of DNWR on wages and employment. In Section III, we describe the SCF data and explain how we construct the wage series. In Section IV, we estimate real wage Phillips curves using both aggregate data for Canada as a whole and disaggregated provincial data. We attempt to reconcile different pieces of evidence by analyzing the evolution of real wages by job seniority in Section IV. We conclude in Section V.

2 Literature Review

In this section, we review some of the recent studies that document asymmetries in the wage change distribution, based on micro level data. While DNWR could clearly be a source of asymmetry in the wage change distribution, other factors like menu costs may also explain the observed asymmetries. We discuss the evidence related to the two hypotheses, we then argue that, from the monetary policy perspective, it is more interesting to examine the impact of DNWR on aggregate wages and consequently on employment. We summarize briefly the current state of the literature that examines this question.

2.1 Asymmetric wage change distribution

The empirical literature using individual level data is expanding very quickly. We will restrict our attention to a few representative papers based on U.S. data, and more recent studies using U.K. Household data and Canadian data. A typical paper in this literature considers the distribution of nominal wage growth in an average year (mostly low inflation year) and highlights the following two visual observations:

- There are relatively few wage cuts, and
- There is a mass point in the wage change distribution at zero.

2.1.1 How frequent are wage cuts?

McLaughlin (1994) documents that nominal wage cuts are not rare in the United States between 1976 and 1986. Using survey data from the Panel Study of Income Dynamics (PSID), he finds that 17 percent of workers with the same employers suffer nominal cuts. Subsequent studies using PSID confirmed these results. In particular, Card and Hyslop (1997) show that in a typical year in the 1980s, 15 to 20 percent of non-job changers had measured nominal wage declines, while Lebow et al. (1995) find a similar proportion of 18 percent, on average, between 1971 and 1988.

Stylized facts from other data sources tend to show similar patterns. Using data from the British Household Panel Study, Smith (2000) finds that, on average, 23 percent of workers suffered nominal wage cuts in their weekly pay over one year span in the 1992-1996 period. In Canada, however, the evidence is less conclusive. The Labour Market Activity Survey (LMAS, 1988-1990) and the Survey of Labour and Income Dynamics (SLID, 1993)

results are similar to the PSID, with the SLID showing a surprisingly large number of wage cuts in 1993. On the other hand, the distribution of wage changes in the wage settlements from the unionized sector's collective bargaining agreements shows virtually no mass below zero wage change.

Akerlof et al. (1996) have argued that the variation in the reported wages in the PSID are an artifact of measurement errors. Although no careful treatment of the measurement error has been conducted on the Canadian data, McLaughlin (1994) and Smith (2000)[2] found that about 5 percentage points of the fraction of wage cuts could be attributed to measurement error, decreasing the frequency of pay cuts to still significant levels of 12 percent in the PSID and 18 percent in the BHPS.

2.1.2 The spike at zero wage change

In all the above mentioned studies, the distribution of nominal wage growth exhibits a large mass point at zero. In the PSID sample, Card and Hyslop (1997) report that the fraction of workers on the same job who experience one-year wage change of zero is 8.3 percent in the 1970s and 16 percent in the 1980s. In the United Kingdom, Smith (2000) shows that this fraction is equal to 9 percent between 1992 and 1996. Crawford and Harrison (1997) report that the fraction of wage freezes is 19.4 percent in the unionized private sector in Canada between 1992 and 1996.

Some institutional factors, unrelated to any underlying rigidities, could however exaggerate the size of the mass point at zero. Long term contracting or rounding could also explain part of the excess mass at zero wage change.

To control for the effect of long term contracts, one can calculate the fraction of workers who received zero wage change over varying horizons. Card and Hyslop show that the mass point at zero in the two-year wage change distribution is reduced to 2.6 percent in the 1970s and 8.1 percent in the 1980s. Over three years, these fractions drop to 1.2 percent and 4.7 percent in the 1970s and 1980s, respectively.[3] In the United Kingdom between 1992 and 1996, Smith (2000) shows that the mass at zero drops to 4 percent for wage growth defined over two years, and to 2.5 percent for wage growth defined over three years. In Canada, Crawford and Harrison report a similar drop in the spike at zero when changing the wage cut definition. The fraction of wage freezes in the unionized private sector between 1992

[2]It is very interesting to note that the BHPS gives interviewees a chance to check their pay slip when reporting their pay, thus reducing substantially the likelihood of measurement errors.

[3]Lebow et al. (1995) perform the same calculation and get slightly smaller numbers.

and 1996 drops to 12.9 percent in the wage change distribution over the life of the contract.[4]

After controlling for rounding problems and measurement errors, Lebow et al. (1995) calculate that almost 40 percent of the spike at zero in the one year wage change distribution is due to rounding, while Smith argues that eliminating measurement error could cut the spike by half. This evidence, however, still indicates a substantial fraction of zero wage changes.

2.2 The source of asymmetries

Since the underlying "true" distribution of wage (or productivity) growth is unobservable, it is difficult to identify the source of distortions to the observed distribution. Two hypotheses, DNWR and menu costs, are usually considered. While both types of rigidities lead to a thinning in the left tail of the distribution and a piling up at zero wage change, menu costs also prevent small positive wage changes from occurring.

If DNWR is only binding to the left of the median wage change in the wage change distribution, and assuming symmetry around the median, then the difference between the two tails of the distribution is important in identifying the source of the rigidity. Alternatively, time variation may help disentangle the effects of DNWR from other sort of institutional factors that might generate asymmetry in the observed wage distribution. For example, if the spike at zero is due to a downward constraint on wages, then, assuming that the shape of the underlying distribution does not vary over time, this constraint should be more binding in low inflation years, and less binding in high inflation years.

Card and Hyslop (1997) use the assumption of symmetry to construct counterfactual distribution of wage growth in the absence of rigidities. Their estimate of the fraction of people affected by downward nominal wage rigidity, adjusted for the effect of menu costs, is around 10 12 percent in the mid-1980s. Their estimates also imply that DNWR may have increased by about 1 percent the average wage growth for hourly-rated non job changers, with smaller effect in the later years of the sample. They conclude that DNWR exerts a small but measurable effect on average wage growth, with a bigger effect in low-inflation years.

Lebow et al. (1995) use the difference between the cumulative frequency of the wage change distribution above twice the median and the cumulative

[4]This fraction is higher in the public sector settlement, where wage freezes are between 56 and 45 percent using different wage change definitions.

frequency of the distribution below zero as an alternative measure of asymmetry. They find that the frequency of wage changes below zero is nearly 4 percentage points lower than expected on the basis of their assumptions. The correlation between this measure of asymmetry and inflation constitute a better test of the DNWR hypothesis. They find that this correlation is only negative and significant for job stayers paid by the hour.

The above evidence could overstate the effect of DNWR if the underlying assumption of a symmetric distribution of wage changes was not satisfied. In fact, McLaughlin (1999) shows that the skewness of wage changes is not limited to censoring of would-be wage cuts and small wage changes. There is evidence of skewness even close to the median. These results present a challenge to the estimates of Lebow et al., and Card and Hyslop.

Intertemporal variation of the wage change distribution provides another way to identify thinning of the distribution below zero. Under the assumption that the shape of the underlying distribution does not change over time, Khan (1997) estimates that, in the PSID sample years 1970 1988, DNWR prevented 9.4 percent of wage earners from receiving nominal wage cuts.[5] However, if the sample in low wage-growth years has lower variance of wage changes, then the tails of the distribution would be thinner even if would-be wage cuts were not censored at zero.[6] To address this issue, McLaughlin (1999) uses a difference-in-difference estimator. His results, still confirm Khan's results, pointing to a thinning of tails below nominal zero of one-third to one-half of would-be cuts.

In summary, both DNWR and the menu costs hypotheses are supported in the data analysis. DNWR clearly act as a constraint on nominal wage changes at the micro level. The next section discusses what is the evidence on how these two hypotheses are reflected in aggregate wages and employment.

2.3 Aggregate effects of DNWR

Only few papers address the macroeconomic implications of nominal wage rigidity on aggregate wages and employment (or unemployment). Applying a hazard model to data for union wage settlements in Canada, Crawford (2000) estimates suggest that the net effect of rigidity on average wage growth, between 1992 1997 is less than 0.2 percent for the unionized private sector. These estimates are significantly lower than those reported in Simpson, Cameron and Hum (1999) for the same data. Using a Tobit model

[5] In contrast, salary earners do not receive pay cuts less frequently than expected.

[6] The changes in the shape of the wage growth distribution is well documented in Crawford (2000).

for wage growth. Simpson et al. estimate that DNWR raised the average wage growth by 0.67 percent between 1993 and 1995. On the other hand, Farès and Hogan (2000) conclude that, consistent with menu costs, nominal rigidities have a symmetric effect on wage changes above and below zero. Overall, they conclude that nominal rigidities result in lower than expected wage changes.[7]

Simpson et al. (1999) also provide some estimates on the effect of pay cut resistance on employment growth and the unemployment rate. They use OLS estimation of employment growth on pay freeze incidences and output growth, in different periods of high and low inflation. Their results indicate that, between 1993 and 1995, DNWR reduced mean employment growth across sectors by more than half. It is worthwhile noting that the wage freeze variable in this regression might be capturing some adverse shocks, particularly since that the output growth estimated effect between 1993 and 1995 is significantly lower than in previous periods. Farès and Hogan (2000) and Faruqui (2000) show that, once adjusted for this endogeneity problem, the effect of wage freezes on employment growth becomes not statistically significant. Using a Tobit specification, Simpson et al. (1999) calculate that the unemployment cost of pay cut resistance exceeds 2 percent throughout the 1990s. One underlying assumption of these estimates is that the variance of the wage growth is time invariant. As discussed above, this assumption could exaggerate the effect of DNWR given the noticeable compression in the wage change distribution in the 1990s, a period of low inflation.

Card and Hyslop use state level average wages and unemployment from 1976 to 1991 to estimate the effect of DNWR on unemployment. They use wage data constructed from the annual March Current Population Survey (CPS) that they adjust to reflect the varying composition of the workforce in each state in different years. They estimate cross-state Phillips curve and find little evidence that the rate of wage adjustment across local markets is faster in a higher-inflation environment. Taken in combination with their micro level findings, they argue that nominal rigidities have a small impact on the aggregate economy.

Overall, the micro level evidence based on the distribution of individual wage changes reveals that although nominal wage cuts are not rare, there is a substantial spike at zero in the distribution of nominal wage changes. Furthermore, there is some evidence that the magnitude of the spike is cor-

[7]Crawford (2000) discussed these results and suggests that different treatment of inflation expectations could reconcile the results from Simpson et al. (1999) and Farès and Hogan (2000).

related with inflation. It is much less clear from the literature, however, that DNWR have significant consequences for aggregate wage and employment (or unemployment) determination. In remainder of the paper, we attempt to fill some of these knowledge gaps by taking a new look at the effect of DNWR on wage and employment determination in periods of low inflation in Canada.

3 Wage Data

3.1 Survey of Consumer Finances

We have assembled sixteen annual microdata files from Statistics Canada's Survey of Consumer Finances (SCF) to construct a consistent wage series over the year 1981 and 1997. The SCF provides large samples of around 40,000 workers for each of these years with the exception of 1983 in which the survey was not conducted.[8] For all the available years, the SCF was conducted in April as a supplement to the Labour Force Survey (LFS) and asked a battery of question about income in the previous year in addition to the usual LFS questions that typically pertain to the reference week.[9]

The SCF contains information on annual income as well as personal and labour-related characteristics of individuals aged 15 years and over. In particular, information is available on wages and salaries and income from self employment in the previous year, labour force status, number of weeks worked last year, full time/part time status last year, number of hours in the reference week, occupation and industry, years of experience and seniority, and educational attainment.[10] Other demographic characteristics, like age, gender, marital status, language spoken, immigration status, geographic location are also available.

The wage measure we use is average weekly earnings expressed in 1991

[8]Public use samples are also available for heads of households and spouses every other year during the 1970s. Data for all workers are only available starting in 1981. The survey was discontinued after 1997.

[9]The reference week is the week immediately preceding the two-week period when the Survey of Consumer Finances is conducted.

[10]One major concern using these data arises from changes in the way educational achievement is classified starting with the 1989 income file. Fortunately, the highest (university degree) and lowest (grade 8 or less) education categories appear to be quite comparable (in terms of sample proportions and average wages) under the two definitions. We exploit this feature of the data later to ensure that our adjusted wage measures are comparable over time.

dollars.[11] For each individual in a given sample year, average weekly earnings are calculated as the ratio of annual wages and salaries, excluding income from self employment and rental property, divided by total weeks worked in that year. We only compute this wage measure for paid workers who report zero net income from self employment to get a cleaner measure of wages for employed workers, since theories of DNWR are not relevant for self employed workers. We also restrict the sample to workers age 20 to 65.

Table 1 presents the distribution of workers across provinces, industries and sectors. About 65 percent of the (weighted) observations are concentrated in Quebec and Ontario, while more than half of the individuals work in the manufacturing, trade and service industries. Only 21 percent of the sample is in the public sector.

The distribution of individual characteristics is presented in Table 2. Beside standard demographic characteristics, the table provides information on full time status and on the distribution of job tenure. Since job tenure is measured at the time of the survey in April, some workers (15 percent of the sample under the "lost their job" heading) report some earnings in the previous year despite the fact that they no longer work at the time of the survey. Another 15 percent of workers have one year or less of tenure at the time of the survey. This indicates a fair amount of labour market turnover.

Table 3 shows the provincial means of log-average weekly earnings for each year. Total average wages do vary substantially across provinces (the last row in the table), with a maximum gap of 26% between Prince Edward Island and British Columbia. By contrast, real wages show relatively little variation over time. In fact, as shown in the last column of Table 3, wages are very stable around their sample average, with the largest difference of 7 percent (drop) between the first year of the sample and the last year.

3.2 Adjusted vs. Unadjusted Wages

There are two potential drawbacks with using average weekly earnings from the SCF as a measure of the wage rate over the business cycle. One first problem is that average weekly earnings may either vary because of changes in the underlying (hourly) wage rate, or because of changes in hours worked per week. Unfortunately, an hourly wage rate cannot be computed directly

[11] Earnings are defined as the sum of wages and salaries from all types of civilian employment. Included are gross cash wages and salaries received in the reference year from all jobs, before deductions for pension funds, hospital insurance, income taxes, Canada Savings Bonds, etc. Tips and net commissions are also included. Taxable allowances and benefits provided by employers are not included.

since the SCF does not provide direct information on the number of hours worked per week in the previous year. Fortunately, several indirect measures of hours worked per year can be used to control for variation in hours. As mentioned earlier, the SCF collects information on hours worked during the reference week and on whether the worker worked full time during the previous year.

We have also computed direct measures of actual hours worked per week by detailed category of worker using the monthly micro data files from LFS from 1981 to 1997. Matching these hours measures to workers in the SCF provides an additional proxy for weekly hours of work in the previous year. Our strategy explained in detail below is to use regression methods to "adjust" average weekly wages for changes in weekly hours of work as proxied by these different measures.

A related issue is that changes in the composition of the workforce may understate the cyclicality of real wages since the skill level of the workforce tends to decrease during expansions and increase during recessions, as younger and less educated workers are the first to lose their jobs in periods of economic downturn (see, for example, Bills, 1987, and Solon, Barsky, and Parker, 1994). As in the case of hours, we control for changes in the composition of the workforce by computing alternative "regression adjusted" measures of the wage rate. More specifically, we use ordinary least squares to estimate the following wage equation:

$$w_{it} = \beta X_{it} + \sum_{t=1}^{16} \delta_t Year_t + \epsilon_{it}, \qquad (1)$$

where w_{it} is log real average weekly earnings of individual i in year t (earnings are deflated by total annual consumer price index); X_{it} includes various observable characteristics such as age, education, sex, marital status, language spoken, tenure, industry dummies, province dummies, full-time dummy, and actual hours of work (in the survey week or for similar workers in the LFS); $Year_t$ is a dummy variable for each year in the sample. The estimated coefficients of the year dummies $\hat{\delta}_t$, $t = 1..16$, can then be interpreted as the regression adjusted measures of the wage rate, i.e. the predicted yearly wage rate of an individual with a fixed set of characteristics.

Figure 1 illustrates the difference between the adjusted and unadjusted wage series. Except for the sharp drop during the 1981 1983 recession, the unadjusted wage shows very little variation throughout the sample horizon. In particular, from 1988 to 1997, this series looks almost flat. By contrast,

movements in various "adjusted" measures of the real wage follow a much more cyclical pattern, with a sharp increase in wages in the late 1980s, and a sharp decrease in the early 1990s. The figure shows three different adjusted measures of the real wages (all series are normalized to zero in 1997 for the sake of comparison). The top line on the graph is the wage adjusted only for changes in human capital (H.C.) and other socio economic characteristics, while the two other wage series are based on models that also control for changes in hours using the hours proxies available in the SCF and the LFS.

The graphs show that using proxies for hours from the SCF or the LFS yield very similar adjusted wage series. The adjusted wage series for which hours are not controlled exhibits more of a downward trend, but its cyclical behaviour is similar to the two other adjusted wage series. In the remainder of the paper, we will use the wage series adjusted for human capital, other socio economic characteristics, and hours as measured in the LFS. Note, however, that the results obtained using the different adjustment schemes are all qualitatively similar.

We use a similar procedure to construct adjusted measures of real wages at the provincial level. More specifically, we estimate a model with a full set of province year interactions:

$$w_{ijt} = \beta X_{ijt} + \sum_{j=1}^{10} \sum_{t=1}^{16} \delta_{jt} Prov_j * Year_t + \epsilon_{ijt}. \qquad (2)$$

Where $Prov_j$, for j=1,..,10, is a set dummy variable for provinces. The estimated province-year effects ($\hat{\delta}_{jt}$) can be interpreted as regression adjusted measures of the wage rate in a province j in year t (i.e. the wage in different provinces and different years for an individual with a specified set of characteristics).

3.3 Comparison with U.S. Wage Series

As an additional check on the quality of our wage series, we compare our results to those obtained using similar data for the United States. In March of every year, the U.S. Bureau of Census conducts an income supplement to the Current Population Survey (CPS) which is very similar to the SCF. Since 1976, the March CPS asks explicitly respondents about their usual weekly hours of work in the previous year. It is thus possible to compute a direct measure of hourly wage rates in the United States by dividing annual wage and salary earnings by total hours of work (product of weeks worked

and hours per week), and to compare this direct measure to the regression adjusted methodology we use for Canada.

Figure 2 shows the unadjusted U.S. series for weekly and hourly wages, as well as the corresponding series adjusted for changes in individual characteristics and hours (in the case of weekly wages).[12] All wage series are pro cyclical though the timing of peaks and troughs in wages tend to slightly precede the peaks and troughs in overall economic activity. Interestingly, the adjusted wage series for hourly wages and weekly wages (series on top of the graph) are very close to each other. This suggest that weekly wages adjusted for the kind of hours measures available in the SCF are a very good proxy for the series based on actual hourly wage rates. Extrapolating these U.S. results for Canada suggest that the time series pattern of the Canadian wage series based on adjusted weekly wages mostly reflect true movements in hourly wages, as opposed to changes in weekly hours of work.

It is also interesting to explicitly compare the Canadian and U.S. wage series. Figure 3 plots the adjusted real weekly wage series (adjusted for individual characteristics and hours of work) for Canada and the United States. The two series are deflated by their own country CPI. In both countries, wages drop in the late 1970s/early 1980s, increase during the recovery of the 1980s, and drop again in the early 1990s. Interestingly, wage changes in the United States tend to precede those in Canada by a few years. For example, real wages drop dramatically between 1979 and 1982 in the United States while this decline only happens between 1981 and 1984 in Canada. In the 1980s, U.S. wages peak between 1986 and 1989 while the peak is only reached in 1989 1991 in Canada. Finally, U.S. real wages fall sharply between 1989 and 1991 while they only start declining (at a slower pace) in Canada after 1990.

One question raised by this figure is whether the very low rates of inflation experienced by Canada in the 1990s prevented real wages to adjust as fast as they should have because of DNWR. Table 4 shows that starting in 1991 1992, the rate of inflation (CPI all items) dropped below 2 percent a year in Canada while it remained around 3 percent in the United States. By contrast, inflation rates in the two countries were roughly comparable during the 1980s. So if low inflation prevented real wages from declining fast enough in Canada relative to the United States, this phenomena should have only occurred after 1991. Figure 3 indicates, however, that real wages

[12]We perform the hours adjustment for the U.S. weekly wage series using the exact same variables as what is available in the SCF, namely full time status in the previous year and hours worked in the reference week.

fell at least as fast in Canada as in the United States after 1991. The big difference between Canada and the United States is that real wages remained constant between 1989 and 1991 in Canada, while they declined sharply in the United States during the same period. Since inflation rates in the two countries were comparable during this period, it is unlikely that DNWR can explain the relative evolution of real wages in the two countries after 1989.

A perhaps more direct way of assessing the role of DNWR in wage determination is to look separately at the evolution of nominal wages and the price level (the two elements used to compute real wages). Figure 4 and 5 plot these two series for Canada and the United States, respectively. The figures show a much sharper break in the trends in these two series after 1991 in Canada than in the United States. In fact, there is almost no nominal wage growth in Canada between 1991 and 1994, which is quite remarkable when compared to other time periods or to the United States. Taken at face value, this suggest DNWR were quite "binding" in Canada in the early 1990s.

In summary, the evidence on the role of DNWR in the relative evolution of real wages in Canada relative to the United States is mixed. While the evolution of nominal wages between 1991 and 1994 suggest that DNWR were quite important, the fact that real wages fell as fast in Canada as in the United States during the same period suggests that DNWR did not prevent real wages from adjusting "fast enough". In light of these ambiguities, we now turn to a more detailed analysis of how DNWR may affect the relationship between real wage changes and economic conditions (unemployment rate).

4 Estimating real wage Phillips curves

As mentioned earlier, a key empirical implication of DNWR is that, in response to a given negative shock, the real wage should decline less in periods of lower than higher inflation. We test this implication by estimating "real wage Phillips curves" that link the unemployment rate to the change in real wages. If DNWR prevents real wages from adjusting (downward) in periods of low inflation, the Phillips curve should be *flatter* in periods of lower inflation than in periods of higher inflation. These models are in the spirit of the traditional Phillips curve approach since *changes* in real wages, as opposed to their level, are expressed as a function of the unemployment rate.[13]

[13] Blanchflower and Oswald (1994) suggest estimating a "wage curve" (wage level as a function of the unemployment rate) instead of a Phillips curve, while Card (1995) and

4.1 Aggregate Phillips curves

Figure 6 plots changes in (adjusted) real wages and the unemployment rate at the national level. Both series have been normalized and the unemployment is plotted on an inverted scale to illustrate the co movements between the two series. The figure indicates that the two series track each other remarkably well. This close link between the two series in confirmed in Table 5 which reports OLS estimates of the Phillips curve. More specifically, column 1 reports estimates from a model in which the unemployment rate is the sole explanatory variable. The dependent variable used in all specifications is the change in adjusted (for individual characteristics and hours) real wages.[14] The estimated effect of the unemployment rate is negative and statistically significant. The estimated coefficient implies that real wages decline by 0.8 percent each time the unemployment rate increases by 1 percentage point. The estimated effect is very similar when a linear time trend is also included in the model (column 2).

A closer look at Figure 6 suggests that the relationship between real wage changes and the unemployment rate may have indeed changed after inflation dropped below 2 percent a year in 1991. More specifically, real wages changes stopped dropping and stabilized around minus one percent a year after 1991 despite the fact that the unemployment rate kept rising between 1991 and 1993. Furthermore, real wage declines in 1992 and 1993 were substantially smaller (around -1 percent) than in the recession of 1981 83 (real wage declines around -3 percent) despite the fact that the unemployment rate was comparable (at around 11 percent) in the two recessions.

This breakdown in the relationship between real wage changes and the unemployment rate after 1991 is partly confirmed in the Phillips curve estimates reported in column 3 of Table 5. The "low inflation regime" is simply captured by a dummy variable equal to one in year 1992 and later, and to zero for earlier periods.[15] If the Phillips curve became flatter in this period, the interaction between this "low inflation regime" dummy and the unemployment rate should be positive and statistically significant. The estimated interaction term reported in column 3 is positive, as expected, but is not significant at standard statistical levels.[16]

Blanchard and Katz (1997) suggest otherwise.

[14]Since the SCF was not conducted for the (income) year 1983, we define the wage change for 1984 as the change between 1982 and 1984 divided by two.

[15]This dummy captures most of the time series variation in inflation which hovered around 4 or 5 percentage points for almost all years until 1991 before declining permanently below 2 percent thereafter.

[16]The dummy for the low inflation regime is also included by itself in the regression

Quantitatively speaking, the estimated interaction term implies that the slope of the Phillips curve is about twice as small during the post-1991 low inflation period than earlier. No clear conclusion can be reached from the aggregate time-series analysis, however, because of the imprecision of the results.

4.2 Provincial Phillips curves

The imprecision of the time-series results may not be surprising since only six yearly observations are available in the "low inflation regime" of the 1990s. Since different provinces experienced quite different economic conditions during the 1990s, this additional cross-provincial variation in unemployment rates (and potentially real wage changes) may help improve the precision of the parameters of interest.

One further concern with the aggregate time-series evidence is that other unmodelled economy wide factors have also changed during this period. For example, inflation expectations may have changed after the Bank of Canada switched to a tighter (and low inflation) monetary policy in the early 1990s. Supply shocks may have also shifted the Phillips curve around during this period

A natural way to control for the economy-wide factors is to turn to cross-provincial analysis that relies on variation in economic conditions across both time and provinces to identify potential changes in the slope of the (provincial) Phillips curve. Since different provinces experience different shocks at different points in time, it is possible, in principle, to identify the connection between (provincial) wage changes and (provincial) unemployment rates while controlling for nation wide factors using unrestricted year effects.

More specifically, we estimate the following type of cross-provincial Phillips curve:

$$\Delta \tilde{w}_{jt} = \alpha(j) + \gamma(t) + \beta_t U_{jt} + \epsilon_{jt}, \qquad (3)$$

where $\tilde{w}_{jt}$ is the adjusted average real wage index for province j at time t, with the first difference taken over time; $\alpha(j)$, for j=1,..,10, is a set of province dummies; $\gamma(t)$, for t=82,..,97, is a set of year dummies; U_{jt} is the measured unemployment rate in province j at time t; ϵ_{jt} represents the residual error term.

since the intercept of the Phillips curve (real wage change when the unemployment rate is zero) will likely be different during low and high inflation periods.

In principle, a separate slope of the Phillips curve (β_t) could be estimated for each year. In practice, we estimate specifications similar to those for the aggregate time-series models in which the provincial unemployment rate is either interacted with the inflation rate or with a dummy variable for the "low inflation regime" to test whether DNWR combined with low inflation have flattened the Phillips curve.

Before going to the regression models, it is useful to look at the main trends in real wages and unemployment rates across provinces. Figure 7 plots the unemployment rate and the change in real wages for the four largest provinces over the 1982-1997 period. The lower panel of Figure 7 shows that, as is well known, the recession of the early 1980s was more pronounced in the West (Alberta and British Columbia) than in Central Canada (Quebec and Ontario). Interestingly, real wages also fell more precipitously in Western than in Central Canada (upper panel of Figure 7). This illustrates a clear tradeoff between the evolution in provincial unemployment rates and changes in real wages, i.e. a cross-provincial Phillips curve.

The regional patterns in the recession of the early 1990s are very different from those of the recession of the early 1980s. Quebec and especially Ontario experienced much steeper increases in unemployment than the Western provinces. Unlike the 1980s, however, there is no clear visual evidence that real wages fell more precipitously in Ontario than in the West. This suggest that DNWR coupled with low inflation may have prevented real wages to adjust as much as they should have in Ontario.[17]

Table 6 shows the OLS estimates of equation 3 using a variety of specifications. In all models we include an unrestricted set of province dummies to absorb permanent differences in wage changes and unemployment rates across provinces. In the first four columns, the slope of the Phillips curve is assumed constant over time. The model in column 1 includes no control for year effects, while column 2 includes a linear trend and column 3 includes a set of unrestricted year effects. The model reported in column 4 includes different linear trends by province in addition to the unrestricted set of year effects (at the national level). In all four cases, the unemployment rate has a negative effect on changes in real wages. The point estimates indicate that a one percentage point increase in the provincial unemployment rate reduces provincial real wage growth by 0.3 to 0.6 percent. The estimated effects are statistically significant for all specifications except the one in column 4.

[17]Some could argue, however, that policies of the provincial government during this period may have also contributed to keeping real wages from falling more.

Columns 5 to 7 report estimates for the same three specifications as in columns 1 to 3 when the provincial unemployment rate is interacted with the dummy variable for low inflation. As expected, the interaction term is estimated much more precisely using cross-provincial variation than when only aggregate variation was used in Table 6. The standard error is around 0.001 as opposed to 0.004 in Table 5. The point estimates of the interaction term are now small and not statistically significant for all the reported models. The same conclusion is reached in column 8 where the actual inflation rate (as opposed to a dummy for low inflation years) is interacted with the unemployment rate. All in all, the cross-provincial estimates do not support the view that the slope of the Phillips curve is flatter in years of very low inflation than in other years.

5 Reconciling the pieces of evidence: who DNWR bind for?

There are some contradictory pieces of evidence regarding the importance of DNWR in the previous sections of the paper. On the one hand, we have shown that there was almost no nominal wage growth in Canada during the 1991-1994 period and that real wages did not fall as fast in this period as in the 1980s recession. On the other hand, our Phillips curves estimates do not suggest that the slope of the Phillips curve became flatter during years of very low inflation than during other years, at it should if DNWR prevented real wages from adjusting enough in the face of negative unemployment rate shocks. Furthermore, real wages fell as fast in Canada as in the United States, where the inflation rate was higher, during the 1991-1994 period.

One possible way of reconciling these seemingly contradictory findings is to exploit the richness of the SCF data to understand better the dynamics of real wage adjustment along the business cycle. As mentioned in the literature survey, theories of DNWR are most relevant for more "stable" workers who are most likely to stay with the same employer. By contrast, DNWR should not prevent employers from hiring new workers at lower nominal wages than they may have done in other circumstances. If the bulk of wage adjustments over the business cycle happen at the entry level, the presence of DNWR may not have much impact on (upward or downward) aggregate wage adjustments.

For example, Beaudry and DiNardo (1991) show that, consistent with implicit wage theory, *real* wages of workers who stay with the same employer are downward rigid. Aggregate real wages only decline during recessions be-

cause of workers who start new jobs. During expansions, real wages may either increase because new workers get higher wages or because workers still with the same employer get pay increases (to prevent other employers from poaching them).[18] Taking Beaudry and DiNardo's results at face value suggest that DNWR should have no effect on aggregate wages and employment. Of course, when inflation gets very close to zero, nominal rigidities are the same thing as real rigidities. They can appear to have an effect to the extent that real rigidities also have an effect.

The SCF data allow us to examine these issues by looking at the evolution of real wages for different levels of job seniority. Figure 8 shows the adjusted wages between 1981 and 1997 for the different levels of seniority available in the SCF. The most noticeable feature of this figure is that real wages of more senior workers are much less cyclical than those of less senior workers. For example, the real wages of workers with 20 years or more of seniority hardly fall at all during the recession of the early 1980s. By contrast, real wages of workers with a year of less of seniority (workers on "new jobs") fell by almost 20 percent during the same period.[19]

Note also that real wages of workers with a year or less of seniority fell by much less in the recession of the 1990s than in the early 1980s. Since DNWR should not play an important role for these workers, this suggests that factors other than DNWR must explain why real wages did not fall as much during the recession of the early 1990s than during the one of the early 1980s. For the most senior workers, real wages appear relatively rigid over the business cycle throughout the 1981-1997 period. The years of very low inflation since 1991 are not different from other years in this regard.

The behaviour of real wages for the different groups may help explain why DNWR may not have much impact on aggregate wages an employment despite the fact they are "binding" in some circumstances. As mentioned earlier, DNWR most likely matter for senior and stable workers who have long term relationships with their employers. For this groups, however, Figure 8 suggests that real wages are quite rigid anyway (for other reasons such as implicit contracts, for example). This means that DNWR matter most for workers whose real wages are relatively inflexible anyway. By contrast, most of the real wage adjustments over the business cycles are accounted by workers on new jobs for which DNWR should not matter very much.

[18] McDonald and Worswick (1999) find similar results for Canada (Beaudry and DiNardo (1991) use U.S. data)

[19] Individuals in the "lost their job" category report earnings during the previous year despite the fact that they were no longer employed at the time of the survey. Their wages can be thought as wages for workers who were about to lose their jobs.

6 Conclusion

One main contribution of this paper is to develop a series of adjusted real wages for Canada from 1981 to 1997. This series is constructed using detailed data from the SCF that allows us to control (adjust) for composition effects over the business cycle. One first finding is that real wages are clearly procyclical in Canada, and that failure to adjust for changes in the composition of the workforce tend to understate the cyclicality of real wages.

We use these wage data to test whether DNWR tend to flatten the relationship between real wages and economic conditions as captured by the unemployment rate. While the aggregate results are indecisive because of small sample sizes, the results based on cross-provincial variation indicate that the slope of this real wage Phillips curve has remained constant over time. These findings suggest that DNWR did not have a significant impact on wage and employment determination during the post-1991 period of very low inflation.

We attempt to reconcile this finding with the rest of the literature that clearly show the existence of DNWR by analyzing the evolution of real wages for different groups of workers. Our results suggest that DNWR bind most for more senior workers who would have relatively rigid real wages even in the absence of DNWR. By contrast, the bulk of real wage adjustments over the business cycle are experienced by new entrants (young workers or workers on new jobs) for which DNWR least likely bind. This may explain why DNWR have little effect of aggregate real wage determination despite the fact that they are a significant phenomena for some groups (older and more senior workers).

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A Tables and Figures

Table 1: **Provincial and Industrial Distribution of the Workforce 1981–1997**

	Sample composition (%)
Province	
Newfoundland	1.91
Prince Edward Island	0.46
Nova Scotia	3.11
New Brunswick	2.61
Quebec	25.37
Ontario	38.44
Manitoba	3.62
Saskatchewan	2.86
Alberta	9.44
British Columbia	12.19
Industry	
Agriculture	1.31
Other primary	2.39
Manufacturing (Non Durables)	9.00
Manufacturing (Durables)	8.67
Construction	5.74
Transportation, Communication	8.02
Wholesale Trade	4.71
Retail Trade	11.77
Finance, Insurance, Real Estate	5.95
Community Services	19.16
Personal Services	7.94
Business and Miscellaneous	7.88
Public Administration	7.47
Sector	
Private	81.35
Public	18.65

Sources: Statistics Canada, Survey of Consumer Finances. Cross-sectional files from 1981 to 1997. No data is available for 1983. Sample size is 623,875.
Notes: The estimated frequency distributions are all weighted.

Table 2: **Distribution of Workers Characteristics, 1981-1997**

	Sample Composition (%)
Age Group	
20 30	32.11
31 40	29.68
41 50	22.91
50 65	15.30
Education	
No schooling or grade 8 or lower	6.64
Grade 9-10	9.27
Grade 11-13 (Did not graduate)	10.73
Grade 11-13 (Graduate)	18.59
Some post-secondary (No Diploma...)	10.59
Post-secondary (Diploma or Certificate)	28.08
University Degree	16.11
Job Tenure	
Less than 7 months	8.89
7 to 12 months	7.73
1 to 5 years	26.28
6 to 10 years	16.25
11 to 20 years	16.92
Over 20 years	8.62
Lost their Job	15.31
Full time Status	
Full Time	83.81
Part Time	16.19
Gender	
Male	53.25
Female	46.75
Marital Status	
Single	24.47
Married	67.61
Other	7.92
Mother Tongue	
English	59.85
French	21.19
Other	18.96

Sources: Statistics Canada, Survey of Consumer Finances. Cross-sectional files from 1981 to 1997. No data is available for 1983. Sample size is 623,875.
Notes: The estimated frequency distributions are all weighted.

Table 3: **Log-Average Real Weekly Earnings By Province**
1981 1997

	Provinces										
Year	Nfld	PEI	NS	NB	QC	ON	MB	SK	AB	BC	Total
1981	1.49	1.30	1.40	1.42	1.56	1.57	1.45	1.51	1.66	1.68	1.50
1982	1.46	1.21	1.38	1.42	1.53	1.50	1.43	1.45	1.65	1.64	1.47
1984	1.38	1.27	1.36	1.43	1.51	1.50	1.47	1.45	1.58	1.53	1.45
1985	1.39	1.31	1.38	1.41	1.50	1.52	1.45	1.43	1.55	1.54	1.45
1986	1.37	1.31	1.36	1.39	1.50	1.56	1.40	1.41	1.57	1.53	1.44
1987	1.39	1.32	1.40	1.35	1.51	1.57	1.40	1.38	1.52	1.52	1.43
1988	1.41	1.28	1.40	1.39	1.48	1.60	1.41	1.42	1.55	1.57	1.45
1989	1.45	1.34	1.42	1.41	1.52	1.58	1.42	1.40	1.53	1.55	1.46
1990	1.39	1.33	1.44	1.40	1.54	1.58	1.40	1.33	1.53	1.58	1.45
1991	1.37	1.31	1.36	1.39	1.50	1.56	1.36	1.33	1.50	1.59	1.43
1992	1.40	1.33	1.34	1.42	1.49	1.59	1.42	1.37	1.51	1.56	1.44
1993	1.37	1.34	1.40	1.40	1.48	1.58	1.40	1.37	1.50	1.52	1.43
1994	1.38	1.35	1.37	1.40	1.50	1.58	1.41	1.39	1.51	1.57	1.44
1995	1.42	1.33	1.35	1.37	1.49	1.54	1.40	1.40	1.47	1.62	1.44
1996	1.41	1.38	1.34	1.39	1.53	1.58	1.42	1.42	1.52	1.58	1.46
1997	1.30	1.31	1.33	1.37	1.51	1.59	1.43	1.38	1.51	1.60	1.43
Total	1.40	1.31	1.38	1.40	1.51	1.56	1.42	1.40	1.54	1.57	1.45

Sources: Statistics Canada, Survey of Consumer Finances. Cross-sectional files from 1981 to 1997. No data is available for 1983. Original Sample size is 623,875.

Notes: Average weekly earnings are calculated by dividing reported wages and salaries (in hundred of dollars) by the number of weeks worked. Individual weights are used to calculate yearly averages. Total Consumer Price Index (CPI=100 in 1991) was used to deflate nominal wages.

Table 4: **The Aggregate Data**

	Canada		United States	
Year	Δp_t	*U.R.*	Δp_t	*U.R.*
1981	11.70	7.58	9.48	7.60
1982	10.26	10.97	6.30	9.70
1984	4.22	11.31	4.22	7.50
1985	3.89	10.68	3.49	7.20
1986	4.09	9.66	1.84	7.00
1987	4.25	8.83	3.58	6.20
1988	3.97	7.77	4.05	5.50
1989	4.88	7.56	4.70	5.30
1990	4.65	8.13	5.26	5.50
1991	5.47	10.33	4.12	6.70
1992	1.48	11.15	2.96	7.40
1993	1.83	11.36	2.94	6.80
1994	0.17	10.38	2.52	6.10
1995	2.14	9.44	2.79	5.60
1996	1.56	9.65	2.91	5.40
1997	1.61	9.12	2.26	4.90

Sources: CANSIM for Canada, BLS for the United States.
Notes: Price changes are calculated as log differences. Annual Changes in Total Consumer price index is our inflation measure.

Table 5: **Estimated Aggregate Phillips Curve**
Sample Years 1982 1997

Dependent Variable: $\Delta\tilde{w}_t$ (change in adj. wage)			
Control Variables:			
Constant	0.077	0.081	0.093
	(0.019)	(0.018)	(0.022)
u_t	-0.008	-0.008	-0.010
(un. rate)	(0.002)	(0.002)	(0.002)
Linear Trend		0.0009	
		(0.0005)	
$Y1992$			-0.037
			(0.049)
$u_t * Y1992$			0.004
			(0.004)
R^2	0.52	0.59	0.55

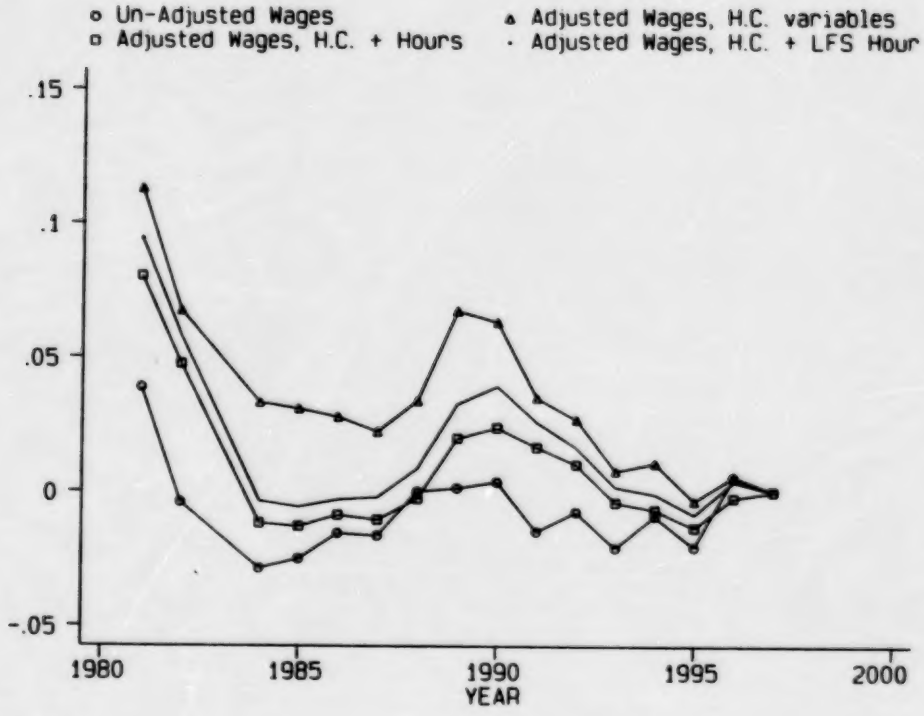
Sources: Statistics Canada, Survey of Consumer Finances, for the wages. CANSIM for prices and aggregate unemployment.
Notes: Standard errors are in parentheses. All regressions are weighted. Annual Changes in log Total Consumer price index is the inflation measure. $Y1992$ is a dummy variable set to one if the year is greater or equal to 1992. For 1984, $\Delta\tilde{w}_{1984} = (\tilde{w}_{1984} - \tilde{w}_{1982})/2$.

Table 6: **Estimated Provincial Phillips Curve**
Sample Years 1982 1997

Dependent Variable: $\Delta\tilde{w}_{jt}$	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Control Variables:								
Constant	0.066	0.070	0.042	0.043	0.067	0.072	0.033	0.051
	(0.013)	(0.013)	(0.019)	(0.023)	(0.014)	(0.014)	(0.022)	(0.020)
u_{jt}	-0.006	-0.006	-0.003	-0.003	-0.006	-0.005	-0.003	-0.005
(un. rate)	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)	(0.002)
Linear Trend		0.0009				0.002		
		(0.0004)				(0.000)		
Year effects	No	No	Yes	Yes	No	No	Yes	Yes
Province Trends	No	No	No	Yes	No	No	No	No
$Y1992$					0.008	-0.010	0.017	
					(0.017)	(0.019)	(0.019)	
$u_{jt} * Y1992$					-0.0003	-0.000	-0.001	
					(0.0017)	(0.001)	(0.001)	
$u_{jt}\Delta p_t$								0.042
								(0.036)
R^2	0.13	0.15	0.19	0.14	0.13	0.15	0.19	0.19

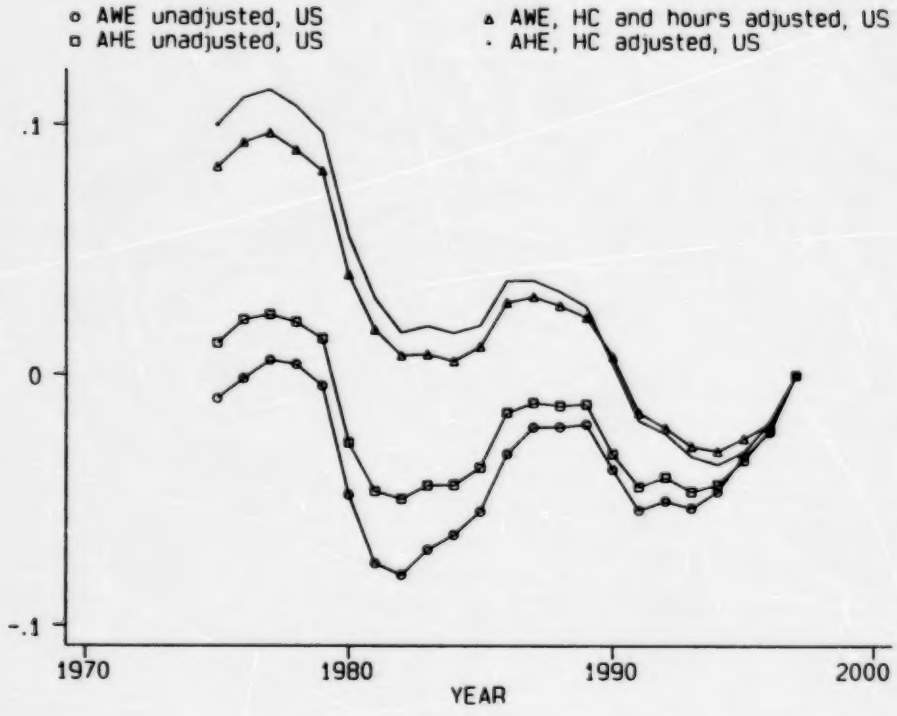
Notes: All regressions are weighted using province weights. Annual Changes in log Total Consumer price index is the inflation measure. All specifications include 10 province dummies. The number of observations is 150. Excluded year is 1997 and excluded province is British Columbia. For 1984, $\Delta p_{1984} = p_{1984} - p_{1983}$ and $\Delta w_{j,1984} = (w_{j,1984} - w_{j,1982})/2$. *Sources:* Statistics Canada, Survey of Consumer Finances, for the wages. CANSIM for prices and provincial unemployment.

Figure 1: Adjusted vs. Unadjusted Wages in Canada



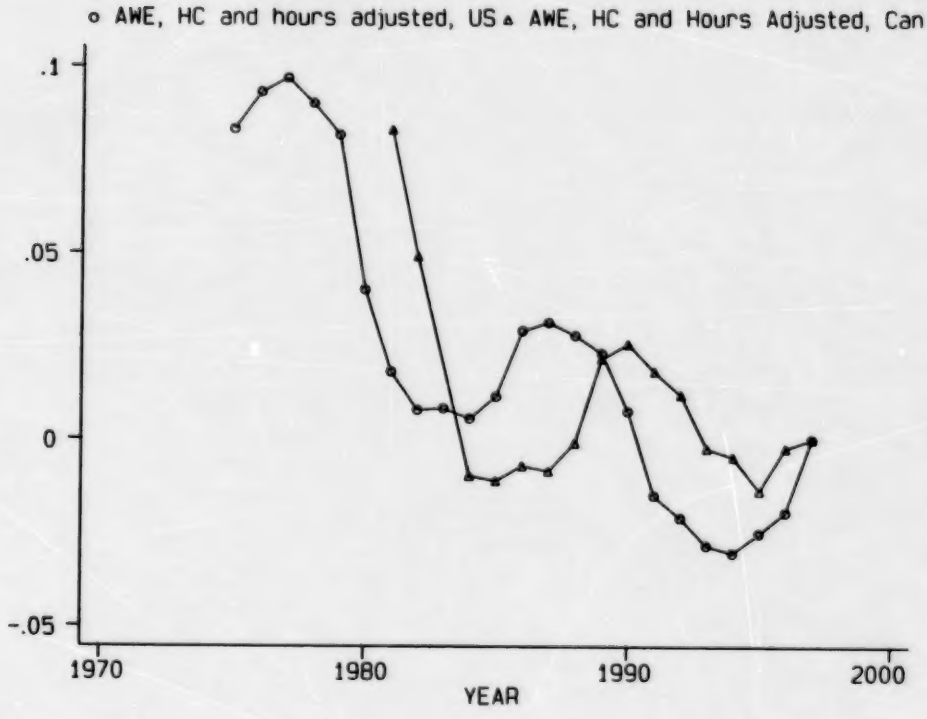
Notes: All wage indices are normalized to zero in 1997.

Figure 2: Adjusted vs. Unadjusted Earnings in the United States



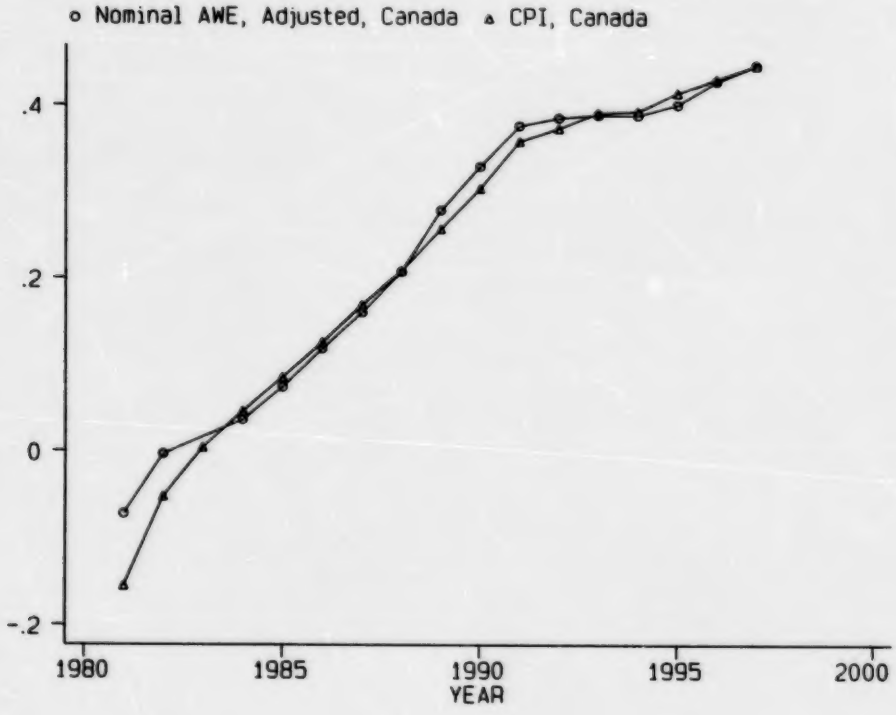
Notes: All wage indices are normalized to zero in 1997.

Figure 3: U.S. and Canadian Adjusted Wages



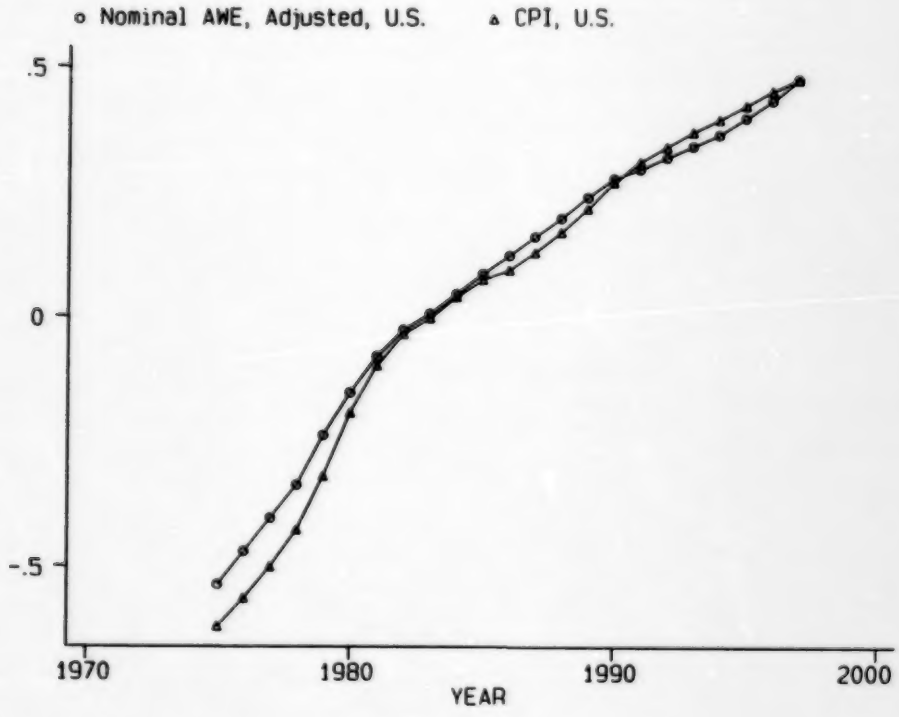
Notes: All wage indices are normalized to zero in 1997.

Figure 4: Nominal Earnings and CPI in Canada



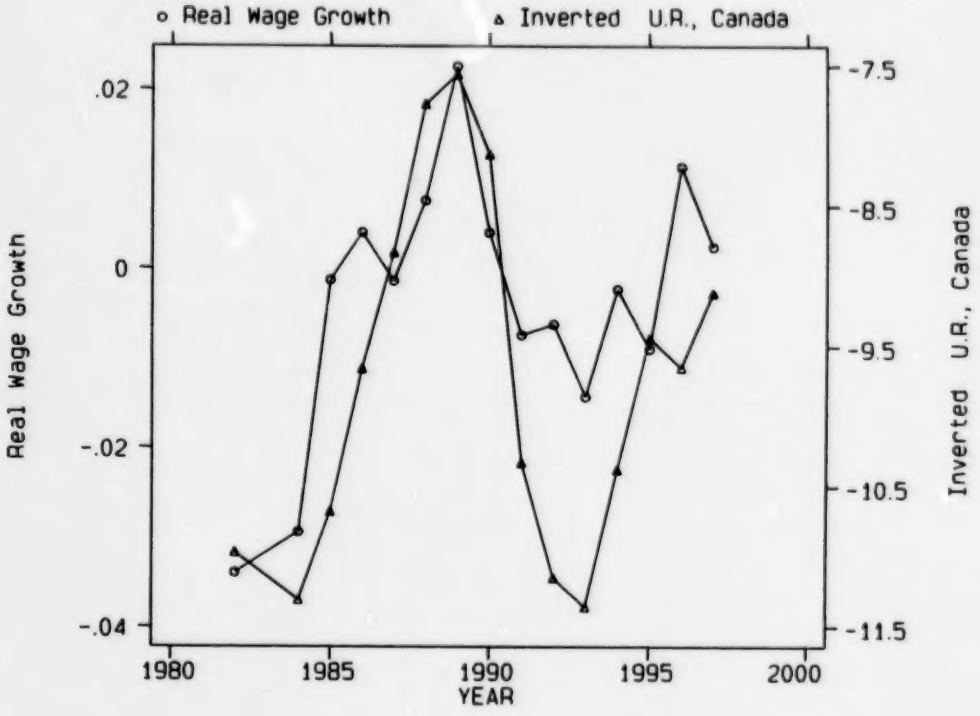
Notes:

Figure 5: Nominal Earnings and CPI in the United States



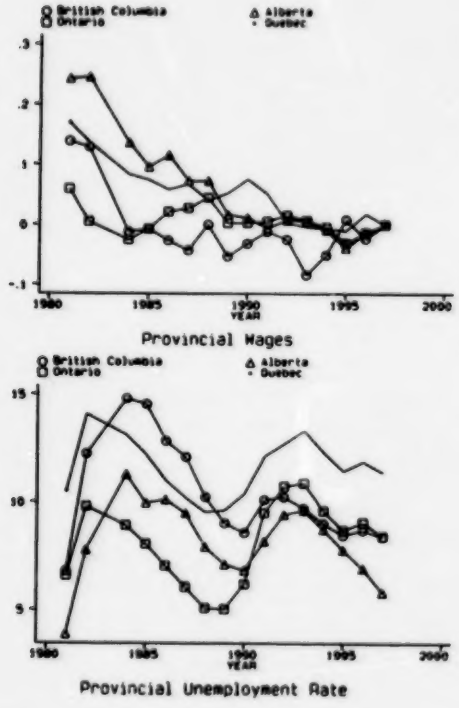
Notes:

Figure 6: AWE growth and Aggregate Unemployment in Canada



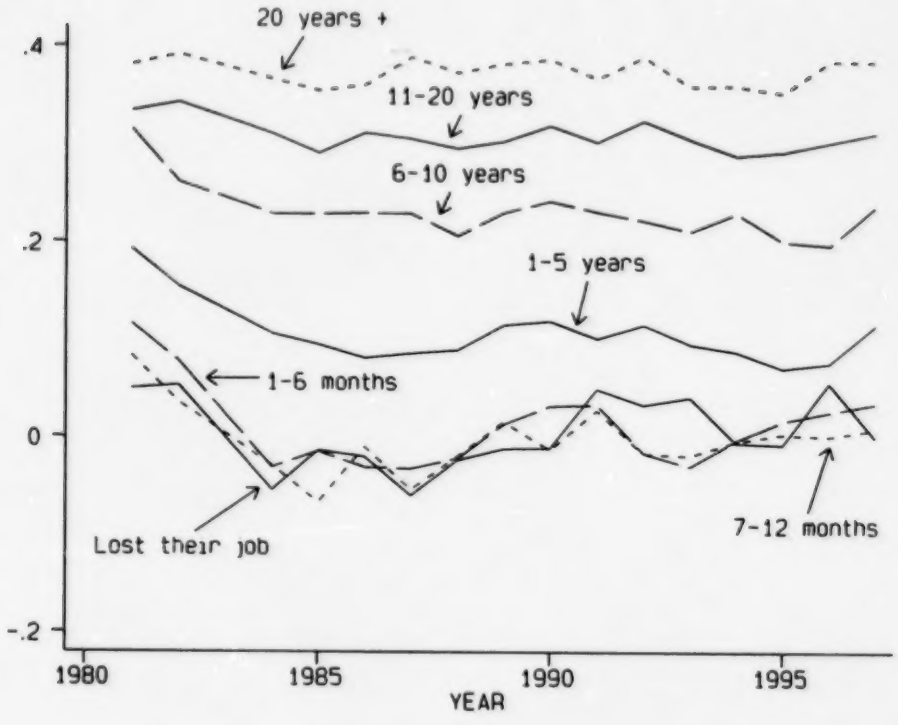
Notes:

Figure 7: Adjusted Provincial Wages and Unemployment Rates



Notes: All wage indices are normalized to zero in 1997.

Figure 8: Adjusted Earnings for Different Job Tenures



Notes: From top to bottom, the lines in the graph correspond to different seniority groups.